

## Assessing French inflation persistence with impulse saturation break tests and automatic general-to-specific modelling

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**Assessing French Inflation Persistence with Impulse Saturation Break Tests and Automatic General-to-Specific Modelling**

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For Peer Review

# Assessing French Inflation Persistence with Impulse Saturation Break Tests and Automatic General-to-Specific Modelling

Carlos Santos (Portuguese Catholic University, Department of Economics and  
Management) and Maria Alberta Oliveira (ISMAI)

## Abstract

This paper has three different motivations. Firstly, we wish to contribute to the debate on whether French inflation has been persistent since the mid-eighties. Empirical evidence in this domain has been mixed. We use the standard method of testing for breaks in the mean of the inflation series to conclude whether possible unit root findings are the result of neglected breaks. Then, we build standard autoregressive representations of inflation, using an automatic general-to-specific approach. We conclude against inflation persistence in the sample period, and the point estimates of persistence we obtain are several percentage points below those achieved with other break tests and model selection methods. Moreover, our final model is congruent. Secondly, we provide the first empirical application of the new impulse saturation break test. The resulting estimates of the break dates are in line with other literature findings and have a sound economic meaning, confirming the good performance the test had revealed in theoretical and simulation studies. Finally, we also illustrate the shortcomings of the Bai-Perron test when applied to a small sample with high serial correlation. Indeed, we show the Bai-Perron break dates' estimates would not allow us to build a congruent autoregressive representation of inflation.

Keywords: Inflation Persistence; Break Tests; Model Selection; General-to-Specific  
*JEL Codes:* E31; E65; C12; C22; C51

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# 1 Introduction

Batini and Nelson (2002) and Batini (2002) distinguish between three different concepts of inflation persistence. The first of these relates to positive serial correlation in inflation, and is the least interesting from the economic point of view. The others refer, respectively, to lags between systematic and unsystematic policy changes and inflation responses. Here the crucial feature is the speed of adjustment of inflation, and that is indeed what one wishes to highlight with the notion of persistence in inflation. Willis (2003) defines persistence as the speed at which inflation returns to its baseline value after a shock. Marques (2004) adapts Willis' definition slightly to argue that persistence is in fact the speed at which inflation converges to an equilibrium after a shock.

The key issue being speed of adjustment, it becomes evident why inflation persistence is a concern among economists and central bankers. As argued by Kool and Lammertsma (2003), a high level of inflation persistence increases the costs of disinflation, and ultimately endangers the disinflation process (namely from the political point of view, due to the social costs associated with the time length of the sacrifice ratio).

In this paper, we shall examine whether or not French inflation has been persistent since the mid-eighties. Following an established tradition in this literature (see, *inter alia*, Bilke (2004a)), we search for a break in the mean of the series prior to testing for persistence. As we shall later argue, there is conflicting empirical evidence in this domain with respect to the post *franc fort* policy (adopted in March, 1983). In particular, different break tests and different modelling strategies have led to the identification of different break dates and to finding different numbers of breaks (see, *inter alia*, Orlandi (2003) and Bilke (2004a)).

Although this paper is only interested in the French case, as there is a vivid debate on whether or not there is a break in French inflation after 1984, this is clearly a small part of a much wider discussion on the linkages between unit roots, structural breaks and inflation. The core of such debate refers to the findings of Pivetta and Reis (2004) for the US economy (see, e.g., Kobayashi (2005)). We shall refer to the debate on the US data in the next section. Here, we shall only reference some results that have appeared in the literature for other countries. Lu and Zheng (2003) conclude that inflation inertia in China is modest. Charezmeh, Hristova and Burridge (2005) find that, out of ninety-three world inflation series, few are deemed to be stationary. Akins and Chan (2004) find that the inflation rate in Canada is stationary around a deterministic trend with two breaks. Finally, Daunfeldt and Luna (2001) conclude in favour of high inflation persistence both in Sweden and in New Zealand in the 1970s and 1980s.

In this paper, we use a new methodology to test for breaks, based on the impulse saturation algorithm (Santos, Hendry and Johansen (2007)). The properties of the impulse saturation break test are explored in Santos and Hendry (2006, 2007). In short, the impulse saturation break test is a procedure to test for multiple breaks at unknown dates, based on a general-to-specific modelling strategy (for results on the consistency of the strategy, see Campos, Hendry and Krolzig (2003)); distributional results on the break date estimators have been developed by Nielsen and Johansen (2007), although more work needs to be done before proper confidence intervals for the break dates can be estimated).

The impulse saturation break test is based on testing the individual significance of each of the impulse indicators in all sample subsets. The subsets are defined in such a way that an indicator for each sample observation is tested. All significant impulses from each sample partition are retained in a union model.

Using this procedure, we establish the existence of three inflation regimes in French Consumer Price Index (CPI) inflation over the 1984Q1 (1<sup>st</sup> Quarter) – 2003Q2 (2<sup>nd</sup> Quarter) period, with break dates in 1993Q2, 1996Q2 and 2001Q2. An economic interpretation of these dates is shown to be possible, contrary to what happens when the Bai and Perron (1998) test is applied to the same data. Using step dummies for the break periods, it is possible to achieve a congruent representation for French CPI inflation. We conclude that there are no signs of persistence in this series.

This paper is organized as follows. Section 2 discusses briefly the economics of inflation persistence. Section 3 presents several scalar measures of persistence. Section 4 overviews the literature on the new impulse saturation break test used in this paper. Section 5 describes the data. Section 6 reports the results of applying the impulse saturation break test and the Bai-Perron test to the data. A comparison with existing results using other break detection methods is outlined. Section 7 discusses modelling this series and persistence assessment. Section 8 concludes.

**2 The Economics of Inflation Persistence**

It is common in the macroeconomics literature on inflation persistence to refer to the models by Taylor (1979, 1980), Calvo (1983) and Fuhrer and Moore (1995). Notwithstanding, the implications of these models are rather different, as far as persistence in inflation is concerned.

Taylor (1979, 1980) considered an economy where wage negotiation is such that half the contracts are revised in one period and the other half in the other period. Hence, his framework is one where wage bargains take place over two periods. In every two

periods all contracts are negotiated. It can be shown that this specification leads to price inertia, but not necessarily to inflation inertia.

Calvo (1983) formalizes a different setting, where it is assumed that firms adjust their prices infrequently. Opportunities to adjust prices arrive as a Poisson process. Hence, the interval between price changes for a given firm is a random variable.

Although the inflation equations in the Calvo and in the Taylor model look similar, Kiley (2000) has highlighted a fundamental difference between the two settings. Whilst in the Taylor model, no wage remains fixed for more than two periods, the probabilistic structure built in the Calvo model allows for many prices to remain unchanged for an indefinite number of periods. Hence, Calvo's model implies a higher degree of inflation persistence than Taylor's model.

Fuhrer and Moore's (1995) model of the inflation process assumes that, at each moment in time, agents take two factors into wage negotiations: on the one hand, they attempt to achieve a current real wage contract equal to the expected average of the real contract over its two periods. On the other, agents may deviate from this expected average taking into account the business cycle. This backward looking component leads to a degree of sluggishness in Fuhrer and Moore's (1995) inflation equation, that amounts to a certain inflation inertia or persistence. In conclusion, from the theoretical point of view, it is important to distinguish whether a model implies price persistence or inflation persistence.

On the empirical side, evidence has been put forth in order to claim that, at least for the US economy, inflation would be highly persistent, approaching a random walk process: Pivetta and Reis (2004) argue they cannot reject the null of a unit root for US inflation. Furthermore, they claim that the conclusions of Cogley and Sargent (2001), with respect to different persistence regimes in the US, are false: not only is inflation persistent, but



also it has been persistent for a long period of time. Stock (2001) also argues in favour of the high persistence of US inflation. On the other hand, Taylor (2000) and Brainard and Perry (2000) claim that inflation persistence has been high, in the US, until the late 70s, when it began to experience a gradual decline.

The study by Pivetta and Reis (2004) has raised serious concerns in other countries, and eventually lead to the joint effort of central banks, the ECB, the Federal Reserve and a number of academics, within the framework of the Inflation Persistence Network (IPN), in order to properly assess persistence in OECD economies. The general conclusion of this plethora of studies was that inflation was not persistent in these countries, and eventual findings of unit roots in inflation series were due to neglected structural breaks (see Perron (1989, 1990) and Hendry and Neale (1991)). Once those breaks were taken into account, most of the conclusions in favour of persistence vanished (see, *inter alia*, Levin and Piger (2004), Bilke (2004a), and Corvoisier and Mojon (2004)). Notwithstanding, the conclusions of these studies seem to be highly dependent on the particular break testing method chosen (namely, break dates estimates and the number of breaks found). The properties of the break tests used (most commonly sequential break tests like Bai and Perron (1998) and Altissimo and Corradi (2003)) are such that possibly make them inadequate for studies with small samples and high serial correlation (see Bilke (2004a) and Vogelsang (1999)). The need to estimate the variance with nonparametric techniques requires a high trimming factor, implying that even fewer observations are available for break date estimation. This, in turn, artificially reduces the number of admissible breaks, as the trimming factor imposes a minimum regime length (see Bai and Perron (2003a) for a discussion). The alternative of lowering the trimming factor, albeit raising the available observations for estimating the break date, is prone to

inducing biased variance estimates, as few observations are used in the kernel, which in turn might lead to biased estimates of the break dates.

In conclusion, the empirical studies do not seem to provide a solid foundation on which to build a claim with respect to inflation persistence, neither in the US nor, more generally, in most OECD economies. More empirical research, using possibly other break detection methods, seem to be fundamental for this debate.

### 3 Measures of persistence

Following Marques (2004), there are two main approaches to assess inflation persistence. One is based on univariate models, generally of the autoregressive form  $AR(p)$ . In these models, shocks to inflation are unidentified and one possible scalar measure for persistence is the sum of the autoregressive coefficients for all included lags (as suggested by Andrews and Chen (1994)). The other is based on structural multivariate models of inflation, with causal factors. In these models, shocks to inflation would not come from the white noise term in the  $AR(p)$ , but rather from the variables that are thought to explain inflation.

Most of the literature, and indeed also this paper, focus on univariate models of inflation. We shall use the sum of the autoregressive coefficients as our measure of persistence (for a critique of other measures, like the largest autoregressive root, the spectrum at zero frequency and the "half life", see Marques (2004)).

### 4 Impulse Saturation Break Tests

A key recent development in testing for breaks at unknown dates, both in the mean and variance, is the result of the impulse saturation algorithm developed by Santos et al.

(2007). Although the baseline Santos et al. (2007) paper does not address the issue of break testing, but rather a model selection problem, the method has been extended to break testing in Santos and Hendry (2006, 2007) and in Santos (2006).

Santos et al. (2007) show that it is possible to include from the outset as many impulse indicator variables as observations in an econometric model (in feasible subsets of, say,  $T/2$ ). The first set of indicators is included for the first half of the sample, and the significant ones are stored. Then, the other half is examined. Under the null hypothesis that no indicator matters, the authors show that, on average, only  $\alpha T$  indicators are retained per regression, where  $\alpha$  is the significance level, matching the binomial argument exactly. Hence, for small  $\alpha$ , there is no evidence of overfitting, in spite of starting the analysis with a very large General Unrestricted Model (GUM). The authors also show that this result is independent of the sample split used (eg. it would also hold for  $T/3$ ,  $T/4$ , etc). The post-selection asymptotic distribution of the mean and variance in a location-scale model with IID errors is derived, and extensive Monte Carlo simulations support the results. Santos and Hendry (2006) show that the procedure can be extended to a number of dynamic models (namely autoregressive processes, both stationary and with unit roots). Nielsen and Johansen (2007) study the autoregressive case further.

Under the alternative, there are breaks in the mean or in the variance. Santos and Hendry (2006, 2007) and Santos (2006) show that the impulse saturation procedure has good power to detect such shifts, both in static and in dynamic models: empirical rejection frequencies of the null are high for the indicators covering the break period, and very low for the remaining indicators; derivation of the non-centralities of the t-tests used for retention/deletion of indicators confirm that those depend on the magnitude of the shift occurred alone. These results were fundamental to a class of other theoretical

developments as, say, the new super exogeneity tests suggested in Hendry and Santos (2007).

Important properties of the new break test are the possibility of using it in models with lagged dependent variables (something that is not always true for the Bai-Perron test, as discussed in Bai and Perron (2003a), in the example using the Garcia and Perron (1996) US real interest rate), and the fact that the power of the test does not depend on the degree of serial correlation in the data. This is a crucial result, as it implies that the impulse saturation test does not need a trimming factor, making it more suitable for small samples. Nonetheless, a minimum regime length still has to be defined in order to avoid confusing breaks with outliers (this shall be discussed in the next section).

In this paper, we provide the first empirical application of the new impulse saturation break test. Results are confronted with those obtained applying the Bai and Perron (1998) test (which is specific-to-general in nature, whilst impulse saturation is clearly general-to-specific). Both sets of conclusions are then confronted with the existing literature results with respect to French inflation persistence.

## 5 Data

In order to assess inflation persistence in France, we have looked at the data used by Levin and Piger (2004).<sup>1</sup> This is quarterly data referring to CPI inflation, comprising the period between the first quarter of 1984 (1984Q1) and the second quarter of 2003 (2003Q2), a total of 78 observations. The choice of this time span is not neutral, as this is precisely the period for which the debate on whether or not there has been persistence in inflation has been most intense. In fact, there seems to be an agreement in the literature as to the existence of a period of high inflation persistence from the mid 1960s

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<sup>1</sup> All data used in Levin and Piger (2004) was collected from the OECD statistical compendium. We are grateful to Jeremy Piger for having sent us this data.

to the early 1980s. However, there is no consensus whether this continued to be the case from the early 1980s onwards or if persistence has declined. Furthermore, Levin and Piger (2004) argue in favour of their sample choice claiming that, avoiding working with data from the 1970s, they would be safeguarding conclusions against the effects of wage and price controls, which would have been common back then.

## 6 Testing for a Break in the Mean with Bai-Perron and Impulse Saturation

### 6.1 Impulse Saturation Break Test Results

As described in section 4, the impulse saturation break test considers adding 78 indicators to the econometric model representing French CPI inflation. We chose to do so in subsets of  $T/2$ , that is, adding first 39 impulse indicators matching each of the first 39 observations and storing the statistically significant ones; then adding the remaining 39 indicators. This is to say that in each of two steps 39 indicators were added to the location-scale model given by (1):

$$\pi_t = \mu + \eta_t \quad (1)$$

where  $\pi_t$  is inflation at time  $t$ ,  $\mu$  is a constant and  $\eta_t$  is an error. There might seem to exist a contradiction between impulse saturating (1) instead of a proper autoregressive representation of  $\pi_t$ , in light of what was said in section 4. Indeed, Santos and Hendry (2006) and Santos (2006) establish that one could impulse saturate the AR representation directly. Nonetheless, by doing so in (1) a direct comparison with the results of the Bai-Perron method is feasible, since Bai and Perron (2003a) acknowledge they need to use (1) due to the absence of lagged dependent variables. Furthermore, Santos (2006) shows through a Monte Carlo study there is no distortion in the impulse

saturation test when (1) is saturated instead of the AR model. Section 7 will shed further light into this by showing that a congruent representation for the inflation process is possible with the impulses retained after saturating (1).

Figure 1 plots  $(1 - p)$  for each of the 78 single impulse indicators, where  $p$  is the  $p$ -value for each indicator. Following Hendry and Krolzig's (2001) remark on the relationship between the choice of the relevant significance level and the sample size, we have decided that given the small sample we had, a significance level  $\alpha = 0.05$  should be used, so that  $\alpha \times T > 3$ .

**<insert Figure 1 about here>**

In Figure 1, the sample periods for which  $(1 - p) > 0.95$  are those for which the impulses were found to be significant in the relevant intermediate regressions, whilst  $(1 - p) < 0.95$  refers to sample periods where the impulses were insignificant.

Following the discussion in section 4, we shall postulate a minimum length of consecutive significant dummies of the same sign and magnitude for a regime to be well defined. On the one hand, comparability of our results with those using Bai and Perron's (1998) test is improved. On the other, and most importantly, time series regimes need to have some minimum length to avoid confusing structural breaks with outliers (see, *inter alia*, Santos, 2006, for a discussion). Nonetheless, there is no reason why the trimming factor should be the same in both methods, given that Bai and Perron's is related to the impact of serial correlation and to the sample size, whilst here it merely reflects a concern with spurious breaks. Bearing this in mind, and given that we are working with quarterly data, we shall postulate 8 periods as the minimum regime length (smaller than the 12 periods we shall postulate for the Bai-Perron test in the next subsection, due to apparent robustness of the impulse saturation test to serial correlation).

Observations above the horizontal line have  $p$ -values lower than 0.05. An inspection of figure 1 reveals the existence of different regimes: one appearing to last from the beginning of the sample until 1993,<sup>2</sup> then a period of overall insignificant dummies until somewhere in 1996, a period of significant dummies from 1996 to 2001, and then again a period of insignificant indicators.

The claims made above, on the basis of the plot, can be made more precise by looking at the  $p$ -values table.<sup>3</sup> Looking at such information it is possible to assert the existence of the four regimes highlighted in table (1).

**<insert table 1 here>**

Therefore, we claim that, when modelling quarterly CPI inflation in France, in the 1984Q1-2003Q2 period, step indicators should be included, matching some of the periods indicated in table (1). There is a clear need to provide a logical link between identification of different regimes on the basis of dummy significance, as in table (1), and the construction of step indicators (which are basically indices (see Hendry and Santos, 2005) where all observations carry equal weights). This mapping of regimes into some indices was possible here after the analysis of the impulses' estimated individual coefficients (see figure 2). The conditions for the mapping to be possible are that, within each regime, all indicators' estimated coefficients have the same sign and similar magnitudes, as discussed in Hendry and Santos (2005) and in Santos (2003). The analysis of figure 2 shows that all indicators' coefficients in the 1984Q1-1993Q1 period have the same sign, and that most estimated coefficients oscillate around a medium

<sup>2</sup> The period of insignificant dummies between 1989Q1 and 1990Q1 is of insufficient length, according to the aforementioned criterion, for a break point and (hence) a change in regime to be considered. Furthermore, the indicator for 1989Q3 had a  $p$ -value of 0.04, yielding one significant dummy in that period.

1991Q2 is an even more clear example of a case where only one indicator deviates from the mainstream of the period in terms of  $p$ -values. It is therefore the case that this cannot be regarded as a change in regime. These examples highlight the need to define a minimum regime length.

<sup>3</sup> Not reproduced here for space considerations, but available upon request.

value of 3. Clearly, there are some lower ones<sup>4</sup>, but we could foresee that from figure 1, given the existence of a few insignificant indicators in the period. Furthermore, there are some higher ones at the beginning of the sample but not in a meaningful number. Hence, we choose to build a step indicator referring to this period (clearly it entails some misspecification in the index, but this has been dealt with, at the econometric theory level, in Hendry and Santos (2005)).

<insert figure 2 here>

Furthermore, the 1996Q2-2001Q1 regime is also well behaved with all indicators with the same sign, and very similar magnitudes. The same happens for the period 2001Q1-2003Q2, although to a less clear extent. We also create these two step dummies.

## 6.2 Bai-Perron Break Test Results

With respect to the Bai and Perron (1998) test, we need to bear in mind the approach used in Bai and Perron (2003a) with respect to testing for breaks in Garcia and Perron's (1996) US real interest rate data (a location-scale model was used to test for a break in the mean, due to the assumption of no lagged dependent variables in the Bai and Perron (2003a) algorithm). The convergence results in Bai and Perron (2003a) crucially require no lagged dependent variables among the regressors. Hence, we will build a model like (1) above to apply the test and estimate the number of breaks in French CPI inflation in the sample period.

Moreover, to apply the Bai-Perron test we need to choose a trimming factor  $\varepsilon = h/T$ , where  $h$  is the minimum regime length, which is dependent on heterogeneity and

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<sup>4</sup> These insignificant indicators can be easily linked to events in French economic policy history, as accounted for in Bilke (2004b). According to Bilke (2004b), one time changes in the VAT rates are good candidates to explain such insignificant dummies: e.g. the insignificant dummy for 1989Q1 is probably due to the simultaneous decreases in the VAT high rate and in the VAT low rate that took place in January 1989; whilst the decrease in the VAT rate for specific items in the CPI in January 1990 could explain the lack of significance of the 1990Q1 dummy.



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correlation of the data, and on the sample size (see discussion in Bai and Perron (2000, 2003a, 2003b)). Bearing in mind the small sample size and the serial correlation of our data, we have chosen  $\varepsilon = 0.15$  as a trimming factor,<sup>5</sup> and allowed for a maximum number of breaks of  $M = 4$  (that is a maximum of five regimes).<sup>6</sup> Hence, each segment had to contain at least 12 observations.

There might be contradictory evidence in the statistics reported in the Bai-Perron method. Such a possibility is acknowledged by the authors, who recommend the use of the  $\sup F_T(l+1|l)$  statistic to decide, when the results of other statistics are less clear.

In the Bai-Perron procedure<sup>7</sup>, one starts by looking at the question of whether or not a break might exist in the sample. For this purpose we report the observed values of the  $\sup F_T(k|0)$ , with  $k$  ranging from 1 to 4, in table (2). We conclude that at least one break exists in the series, even when the stringent 0.01 significance level is used. This is confirmed by the Dmax and WDmax statistics, as can be seen in table (3).

It is then legitimate to say, contrary to what happened in the Garcia and Perron's (1996) example, used in Bai and Perron (2003a), that the Dmax, the WDmax and the  $\sup F_T(k|0)$  statistics corroborate each others. There is sufficient empirical evidence to conclude, at the significance level  $\alpha = 0.01$ , that there is at least 1 break in the sample.

**<insert table 2 about here>**

**<insert table 3 about here>**

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<sup>5</sup> This is clearly greater than  $\varepsilon = 0.05$ , the value suggested in Bai and Perron (1998). Our choice reflects precisely the serial correlation in the data: allowing for regimes with more observations in order to improve the estimate of the variance.

<sup>6</sup> Bai and Perron (2003a) recommend the default use of  $M = 5$ . However, given the small sample size we are dealing with, this would force us to reduce the trimming factor, obtaining more imprecise nonparametric variance estimates.

<sup>7</sup> The Ox codes used in this paper were written with the invaluable help of Jack Luchetti and are available upon request.

The immediate issue would then be: what would be that break date if one was to consider simply the point at which the sum of squared residuals is minimized? The answer we get from the Ox routine implementing the Bai-Perron test is 1993Q2.

We then turn to the issue of how many breaks exist in this sample. We perform the sequential  $\sup F_T(l+1|l)$  test, for  $l$  ranging from 1 to 3. Results are reported in table (4).

**<insert table 4 here>**

Just considering the break dates, for a moment, and ignoring the assessment of their significance, confrontation of the results in table (4) with those in table (1) leads us to notice that, with the exception of 1987Q1, identified by the Bai and Perron (1998) procedure and not by the impulse saturation breakpoint test, the break dates of 1993Q2 and 1996Q2 are provided by both methods. Furthermore, the estimates for the last break point are fairly close: the Bai and Perron (1998) test suggests 2000Q2, whilst in table (1) we get an estimate of 2001Q2. To obtain a better overall assessment of this comparison, we provide the Bai and Perron (1998) estimated confidence intervals for the break dates. Table (5) highlights the five regimes suggested with the Bai-Perron method, whilst table (6) provides the confidence intervals for each of the estimated break dates. In table (6), we report results both for a 90% confidence level and for a 95% confidence level.

**<insert table 5 here>**

**<insert table 6 here>**

The confidence intervals are not symmetric around the break date point estimate, but that was to be expected, since it was suggested in Bai and Perron (2003a) as a property of their method.

From table (6), the first break date is imprecisely estimated: the confidence interval is wide at both confidence levels considered. The break date of 1993Q2 is estimated with high precision, as the confidence intervals are narrow. The same comment broadly applies to 1996Q2. On the other hand, the final break date has wider confidence intervals, although not as wide as the ones in the first break date picked up. It is interesting to notice that the break dates identified with higher precision are precisely the ones also picked up by the impulse saturation method.

In spite of the fact that the minimization of sums of squared residuals in the Bai-Perron method would lead to similar estimates to the impulse saturation break test, some of the break dates in Bai-Perron are not statistically significant when their specific-to-general break testing device is applied, by means of using the  $\sup F_T(l+1|l)$  statistic, for  $l$  ranging from 1 to 3.  $\sup F_T(2|1)$  is significant, even at the 0.01 significance level. So there is statistical evidence to confirm the existence of at least two breaks. However,  $\sup F_T(3|2)$  is smaller than any critical value at any significance level considered, leading us not to reject the hypothesis of 2 breaks versus the alternative of 3 breaks. The specificic-to-general nature of the procedure leads us to stop the search here. In conclusion, we will claim that the Bai-Perron method suggests two break dates: 1987Q1 and 1993Q2, identifying three regimes in the data.

**6.3 Comparison with other studies**

We shall refer to other empirical studies that have looked at breaks in French CPI inflation in the 1980s and 1990s. Using a test for a single break date, Gadzinski and Orlandi (2003) have concluded that there exists one break in the 1990s: in 1992Q2. In a different approach, using Bai and Perron’s (1998) test but with a different data set,

Rapach and Wohar (2002) find a break within the period between the beginning of 1991 and the end of 1993. Benati (2003) finds a break in the period between 1991 and 1995, using Andrews and Ploberger (1994). Using the Bai-Perron test, Benati concludes for a break between 1988 and 1992. Using a Bayesian approach to break testing, Levin and Piger (2004) find one break only at 1993Q2 (the only period where Bayes Factor is  $\geq 1$ ). Furthermore, we are unaware of any studies of breaks in French CPI inflation in the eighties that would conclude for a break in 1987Q1. In fact, most studies comprise the entire 1980s and hence typically find a break date in the early-mid sample. So the issue of the first regime remains open to dispute.

Given this overview of the literature, it is fair to say that the break date in the early 1990s (1993Q2 both in the impulse saturation break test and in Bai-Perron) is in line with the mainstream results found by Rapach and Wohar (2002), Benati (2003) and Levin and Piger (2004).

The economic interpretation of the breaks we found is easy even for the second half of the sample (the same being said for Bai-Perron break dates estimates). The 1996Q3-2001Q1 regime experiences the lowest average inflation rate across the sample. This is due to anticipated effects of the EMU creation, namely convergence of interest rates and inflation rates with the German ones. Moreover, the higher average inflation after 2001Q2 might reflect the heavy depreciation of the EURO-USD exchange rate experienced in the early stages of the EURO.

Adding to this, finding a break in 1993Q2 strikes us as being reasonable as the early 1990s were an important landmark for many countries as far as a decrease in mean inflation is concerned. Corvoisier and Mojon (2004) put forth two main reasons for this: the spread of inflation targeting and, for the relevant countries, the beginning of the

nominal convergence process foreseen in the Maastricht treaty. The authors claim this break in the early 90s is almost a stylized fact across OECD economies.

In conclusion, results seem to point in the direction that the impulse saturation break test is leading to similar conclusions as the mainstream literature, while, at the same time, it is providing economically meaningful break dates.

## 7 Modelling Inflation Dynamics

### 7.1 Using Impulse Saturation

Following the branch of work in the inflation persistence literature that we are trying to pursue, we shall now progress to build a univariate inflation model of the autoregressive type. Our novelty relative to the previous existing papers will be to use an automatic general-to-specific (GETS) strategy. In fact, we use PcGets 1.15<sup>8</sup> (Hendry and Krolzig, 2003) to build a congruent model of inflation, where the GUM contains eight lags of inflation, and the three step dummies we have come to conclude from the impulse saturation breakpoint procedure that should be relevant. In the final model, only the constant, one and three lags of inflation and two of the step dummies were retained (step1, corresponding to the 1984Q1-1993Q1 period; step2, corresponding to the 1996Q2-2001Q1 period):

$$\hat{\pi}_t = 0.01 + 0.33\pi_{t-1} + 0.18\pi_{t-3} + 0.006step1_t - 0.002step2_t \quad (2)$$

(0.002)    (0.11)            (0.09)            (0.002)            (0.001)

Misspecification tests are reported in table (7)<sup>9</sup>.

<sup>8</sup> PcGets is an Ox (Doornik, 2001) package, designed to implement automatic general-to-specific model selection (see Hendry and Krolzig, 2001).

<sup>9</sup> In table 7, AR stands for the Breusch-Godfrey autocorrelation test (Breusch (1978) and Godfrey (1978)), ARCH stands for Engle's (1982) test and RESET stands for Ramsey's (1969) test for mis-specification

<insert table 7 here>

From the output of the test summary reported in table (7), a congruent representation for inflation dynamics over the sample period has been achieved. Indeed, only the RESET test would pose some problems, but these would vanish if one was willing to work with a 0.05 significance level. Furthermore, given that in this class of problems, inflation is modelled in a univariate setting (without any structural causal variables), it does not seem surprising that any eventual problem should come through the RESET test. We have not reported here the conclusions for Hansen's (1992) parameter stability test, for space considerations. Nonetheless, joint stability is not rejected, neither is individual stability for each of the regressors' coefficients.

The estimated equation residual standard error is  $\hat{\sigma} = 0.0052$ , whilst the reported  $\log Likelihood = 271.836$ .

Table (8) reports a final check we have conducted on the PcGets selected model: a Lagrange Multiplier (LM) test for the excluded variables. The conclusion is that none should have been included.

<insert table 8>

Finally, figure 3 suggests that the residuals from model (2) are displaying a white noise type of behaviour, whilst figure 4 shows the proximity between actual and fitted values.

<insert figure 3 about here>

<insert figure 4 about here>

In conclusion, the impulse saturation approach has provided us with relevant step dummies that once included in the GUM for the inflation equation, allowed the

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errors. The normality test used is the one suggested by Hansen and Doornik (1994), whilst White's (1980) test is used to check homoscedasticity.

construction of a final congruent model, in a general-to-specific way. It should be added that our point estimate for persistence, over the sample period, in French CPI inflation is of about 0.5. This is lower than the value Levin and Piger (2004) found for the same period (0.77), when a Bayesian break detection procedure was employed and a single break date was used (1993Q2). The inclusion of the step indicators is crucial to rule out persistence. In fact, if (2) was to be estimated without the step dummies but including the relevant lags and the constant, the Likelihood Ratio (LR) test for the null of the existence of a unit root would yield an observed statistic of 6.197, which would be smaller than the relevant 95% quantile of the Dickey-Fuller type II distribution:  $\cong 9.1$  – hence not rejecting the null of a unit root (see Perron (1989, 1990) and Hendry and Neale (1991)).

7.2 Using the Bai-Perron results

If we are to consider the estimation results using the step dummies obtained from the Bai-Perron test (s1: 1984Q1-1987Q1; s2: 1987Q2-1993Q2), we select the final model:

$$\hat{\pi}_t = 0.014 + 0.54 \pi_{t-1} + 0.007 s_1 + 0.007 s_2 \tag{3}$$

$(0.087)$

$(0.11)$

$(0.002)$

$(0.002)$

Point estimate of persistence would be of 0.54, which is not too different from the one obtained with impulse saturation. However, model (3) does only include one lag of inflation, and both step dummies are kept by PcGets. Table (9) reports the misspecification test results.

<insert table 9 here>

It is clear from table (9) that the final model selected using the Bai-Perron suggested step dummies is not congruent as there are several misspecification test rejections at a

significance level  $\alpha = 0.05$  (AR 1-5 and RESET). The conclusion seems to be that although we achieve a congruent representation for inflation dynamics in France when using the impulse saturation break detection device, this does not happen when the Bai-Perron break date estimates are used in this setting.

## 8. Conclusion

In this paper we have developed a first empirical application of the new impulse saturation break test. We have concluded, that for this sample, it performs better than the Bai and Perron (1998) test. Problems with the Bai-Perron test had already been anticipated in a few studies (e.g. Prodan, 2003), mainly in finite samples with high serial correlation.

We have used the impulse saturation break test in the context of searching for a break in the mean of French CPI inflation series. Finding such breaks is shown to be fundamental to preclude spurious unit root findings. This result is clearly in line with other literature claims that have not used the same break testing procedure, nor a GETS modelling strategy.

Using the break dates suggested by the new test we are capable of finding a congruent representation for inflation dynamics in France, over the sample period. The same is not true if the Bai-Perron break dates' estimates were to be used.



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TABLES

Period	
1984Q1-1993Q1	significant
1993Q2-1996Q1	not significant
1996Q2-2001Q1	significant
2001Q2-2003Q2	not significant

Table 1: French CPI Inflation Regimes (Levin and Piger’s (2004) data) – Impulse saturation test

$\sup F_T(k 0)$	observed	$c_{\alpha=0.01}$
$\sup F_T(1 0)$	16.494	13
$\sup F_T(2 0)$	30.515	9.36
$\sup F_T(3 0)$	31.514	7.6
$\sup F_T(4 0)$	74.759	6.19

Table 2: Bai-Perron  $\sup F_T(k|0)$  test for 1 break: observed test statistic and 1% critical values

Test	observed	$c_{\alpha=0.01}$
Dmax	74.7594	12.37
WDmax	148.43	13.83

Table 3: Further evidence on one break –: Dmax and WDmax observed statistics and 1% critical values

$\sup F_T(l+1 l)$	Observed	$c_{\alpha=0.01}$	$c_{\alpha=0.025}$	$c_{\alpha=0.05}$	$c_{\alpha=0.1}$	Breakpoints
$\sup F_T(2 1)$	21.3884	13	10.18	8.58	7.04	1993Q2 1987Q1
$\sup F_T(3 2)$	5.3889	13.9	11.86	10.13	8.51	1993Q2 1987Q1 1996:Q2
$\sup F_T(4 3)$	10.7628	14.8	12.66	11.14	9.41	1993Q2 1987Q1 1996Q2 2000:Q2

Table 4:  $\sup F_T(l+1|l)$  using global optimizers under the null –: observed test statistics; 1%, 2.5%, 5% and 10% critical values; estimated break dates

Period
1984Q1-1987Q1
1987Q2-1993Q2
1993Q3-1996Q2
1996Q3-2000Q2
2000Q3-2003Q2

Table 5: Regimes Identified in Bai-Perron (Levin and Piger's (2004) data)

C.I. for break dates 90%	C.I. for break dates 95%
1986:3-2000:4	1986:3-2000:4
1992:3-1993:4	1992:2-1994:1
1995:4-1996:4	1995:3-1997:1
1998:2-2000:3	1997:2-2000:4

Table 6: Confidence Intervals (C.I.) for break dates with Bai-Perron testing when M=4

Test	observed	<i>p</i> -value
AR 1-5	0.761	0.582
ARCH 1-4	0.151	0.962
Normality	2.16	0.34
hetero	1.07	0.391
RESET	3.08	0.084

Table 7: Misspecification tests for model (2): observed statistics and *p*-values

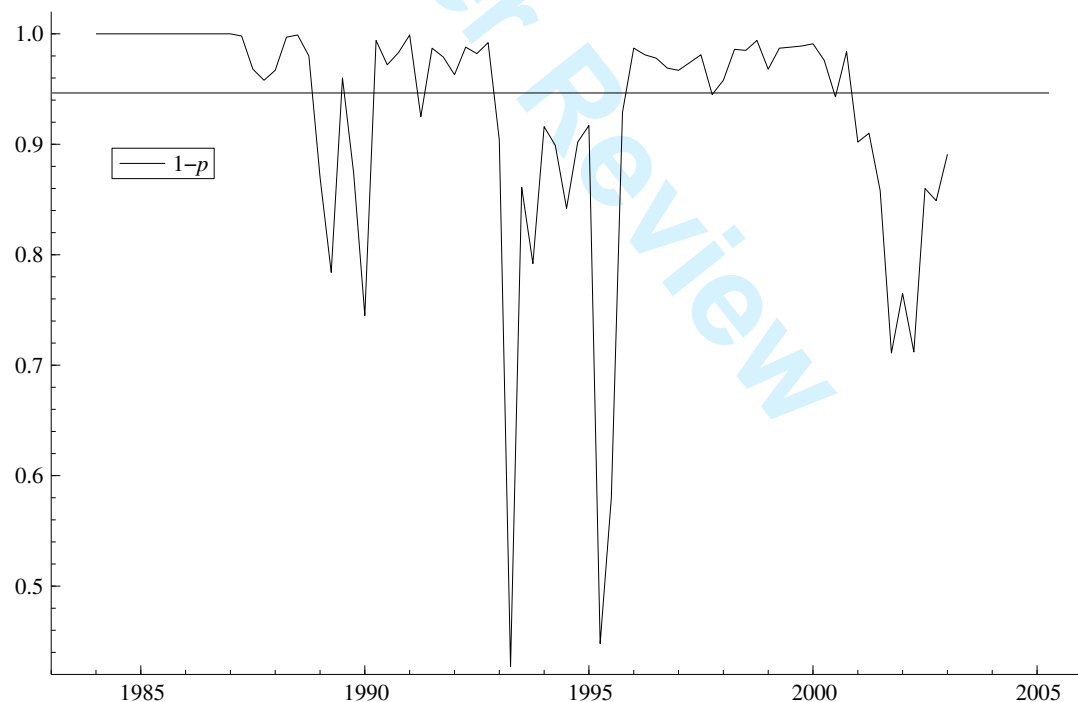
Variable	<i>p</i> -value
inflation_2	0.8082
inflation_4	0.2538
inflation_5	0.3712
inflation_6	0.1679
inflation_7	0.8522
inflation_8	0.833
step3	0.474

Table 8: *LM* test for omitted lags or step indicators in model (2): observed *p*-values

Test	observed	<i>p</i> -value
AR 1-5	2.6074	0.0441
ARCH 1-4	0.477	0.752
Normality	2.528	0.283
hetero	0.389	0.812
RESET	6.88	0.011

Table 9: Misspecification tests for model (3): observed statistics and *p*-values

## Figures

Figure 1:  $(1-p)$  for the 78 indicators; French Quarterly CPI Inflation (Levin and Piger's (2004) data



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Figure 2: Plot of Estimated Coefficients for the 78 indicators: impulse saturation of model (1)

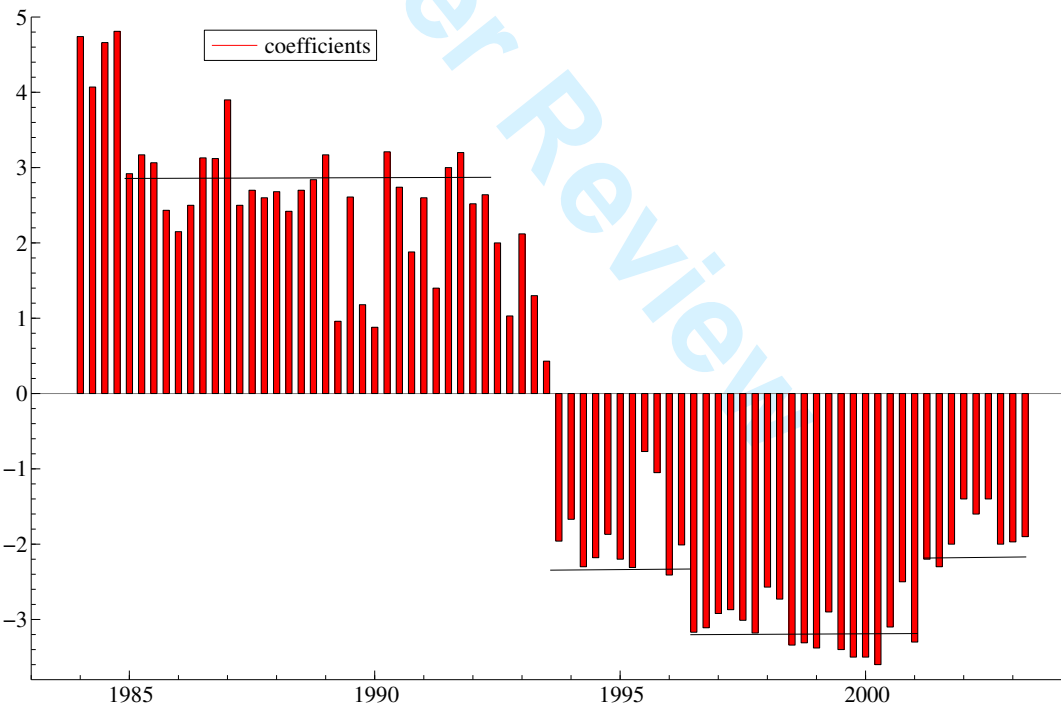


Figure 3: ACF and PACF for residuals from model (2)

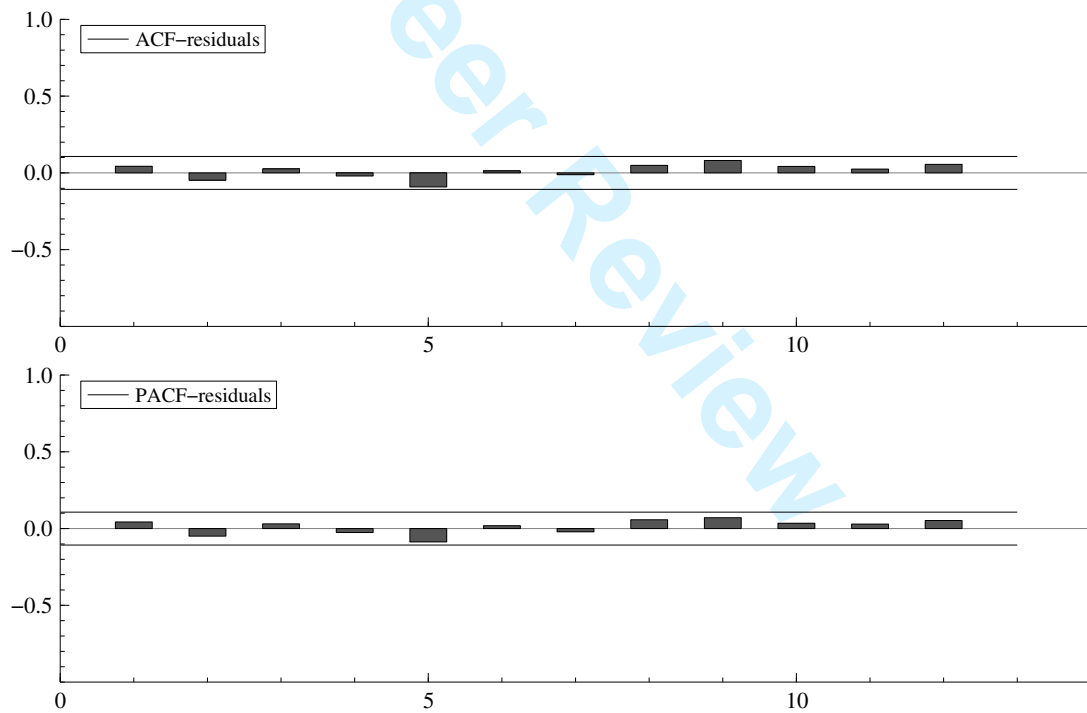


Figure 4: Actual and fitted values from model (2)

